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A Note on Workers' Willingness to Pay
for Nonwage Job Attributes and Labor Mobility

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**A Note on Workers' Willingness to Pay for Nonwage Job Attributes
and Labor Mobility**

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A Note on Workers' Willingness to Pay for Nonwage Job Attributes and Labor Mobility

Abstract. Estimates of workers' willingness to pay for nonwage job attributes (e.g., the risk of injury) are usually based on hedonic wage methods. In this study, workers' marginal willingness to pay for nonwage job attributes is derived from an analysis of job quitting behavior employing discrete choice models. Empirical estimates of the workers' marginal willingness to pay for nonwage job attributes are obtained in two ways. First, estimates are obtained by re-interpreting empirical studies of workers' on-the-job mobility behavior in the US. Second, estimates for workers' willingness to pay for commuting are derived based on an analysis of job quitting behavior in the Netherlands.

1. Introduction

Economists have long been interested in how the theory of compensating wage differentials might explain the existence of wage differences in the labor market. One of the attractions of this theory is that it allows for the estimation of workers' marginal willingness to pay (MWP) for job attributes (Rosen, 1974, 1986), usually referred to as the hedonic wage approach. Although many studies based on this approach have shown that non-wage differences between jobs can be significant to workers, the general conclusion is that non-wage differences between jobs are not very important to workers (see Brown, 1980; McCue and Reed, 1996).

The theory of compensating wage differentials assumes that workers have

complete information in a static environment. This assumption is inconsistent with job search theory, one of the main theoretical frameworks to analyse labour market outcomes. Search theory assumes that workers have incomplete information in a dynamic environment, so workers move from one job to another. Using job search theory, Hwang et al. (1993) have demonstrated that the estimates for the marginal willingness to pay for a job attribute using the hedonic wage approach are likely to be biased *downwards* if it is not acknowledged that a job is a search good and a result of a match between an employer and a worker.¹ It would, therefore, be useful to come up with alternative approaches to estimating the marginal willingness to pay for a job attribute that explicitly recognizes that a job is a search good.

Recently, a number of studies have estimated the MWP for job attributes using data on *job moving* behavior and compared the MWP estimates with conventional estimates (Herzog and Schlottmann, 1990; Gronberg and Reed, 1994; and Van Ommeren et al. 2000).² In line with the reasoning of Hwang et al. (1993), these studies point to

¹ Hwang et al. (1993) reason that hedonic wage approaches are likely to be biased downwards, since the costs of not filling a vacancy are higher for organisations with higher profits. So, these organisations will offer higher wages *and* more attractive job attributes to job seekers to reduce vacancy durations compared to less profitable organisations. As a result, employees working for organisations with higher profits receive higher wages, along with job attributes that are *more* attractive than would be predicted by the theory of compensating wage differentials.

² The notion that job exits are informative on the workers' willingness to pay for a job attribute has some history (see Herzog and Schlottmann, 1990; Van Ommeren et al. 2000). Bartel (1982) argues that the effects of the wage and non-wage characteristics are

considerably higher estimates than those based on conventional hedonic wage methods. Herzog and Schlottmann (1990) and Gronberg and Reed (1994) reported higher estimates for the willingness to pay to avoid job-induced risk. Van Ommeren et al. (2000) reported higher estimates for the MWP to avoid commuting. The view that conventional approaches underestimate the MWP for job attributes is further supported by the study of McCue and Reed (1996), who examined self-reported data on the workers' willingness to pay for job attributes. They concluded that "workers' valuations of nonpecuniary dimensions of work are substantially larger than previous research has indicated".

The current study adapts Gronberg and Reed's (1994) dynamic approach to valuation of job attributes. In their paper, a valuation approach is introduced which has two distinct characteristics: it is based on a structural job search model, so explicitly acknowledges that jobs are search goods, and it requires information on job moving hazard rates. The job moving hazard rate is the instantaneous rate of leaving a job per unit time and can be estimated given job duration data (Kiefer, 1988). Gronberg and Reed (1994) apply their approach estimating a range of job hazard models.

The starting point of this paper is to show that the method of estimating MWP

informative on the "relative values" (or "relative importance") of the job characteristics for the worker. Herzog and Schlottmann (1990) estimate the effects of the wage and the risk in the workplace on the likelihood that the worker switches to another industry, and assess the willingness to pay for risk reduction. They also argue that the latter may differ from the hedonic (market) price of risk if the labor market is imperfect, however, they do not examine a formal behavioural model for the switching rate. Bartik, Butler and Liu (1992) provide a similar analysis for the housing market, estimating the willingness to pay for neighborhood amenities from residential mobility behavior.

for job attributes as introduced by Gronberg and Reed (1994) is not restricted to the use of hazard models, but **can** be extended to the analysis of job mobility using discrete choice models. This gives the opportunity of estimating the **MWP** for job attributes through an alternative **dynamic** approach based on search theory.

In the present paper, MWP estimates are obtained by re-interpreting three previously published empirical studies (Viscusi, 1980; Robinson, 1990; Gruber and Madrian, 1994) that have estimated the relationship between job mobility and job attributes (risk of **injury**, risk of **cancer**, working conditions, promotion opportunities) by **means** of discrete choice models. The aim of these studies was to obtain information on the relationship between job mobility behavior and a number of **nonwage** job attributes, but did not discuss the implied monetary value of these attributes. The present paper **contributes** to the literature by showing that these types of studies contain valuable information on the marginal willingness to **pay** for **nonwage** attributes.

Further, in the **current** paper, the MWP for commuting is estimated using a discrete choice model, and is compared to similar estimates based on a hazard model. We **find** that the **MWP** for commuting estimates based on a discrete choice model are close to those based on a hazard **rate** model.

The outline of the paper is as follows. In **section 2**, theory and estimation method are introduced. In **section 3**, estimates for the marginal willingness to **pay** for job attributes are provided. **Section 4** concludes.

2. The marginal willingness to pay: theory and estimation method

2.1.Theory

In the present paper, we consider a slight extension of a standard on-the-job search model, which excludes nonwage job attributes. Mathematical proofs are omitted and can be found in Gronberg and Reed (1994) and Van Ommeren et al. (2000), respectively.³

The point of departure is an employed individual. This individual derives utility from wage w and from a number of job attributes. For simplicity of notation, we suppose that the individual derives utility from only one nonwage job attribute x .

The person searches in the labor market with a certain effort. Search costs are increasing and convex in search effort. The job arrival rate is increasing and concave in search effort. The effect of the search costs on the instantaneous utility function is additive (see Van Ommeren et al., 2000). Job attribute offers are drawn randomly from a given distribution. The involuntary separation rate (due to firing) is given for the workers. Pooling of offers is not allowed: job offers are either refused or accepted before other offers arrive. The individual is assumed to maximize lifetime utility and the environment is stationary. Given these assumptions, the optimal search effort and the reservation wages can be determined (see Albrecht et al., 1991).⁴ In the present context, the marginal willingness to pay for an attribute can be defined as the marginal lifetime utility of the attribute over the marginal lifetime utility of the wage (see Van Ommeren et al, 2000).

³ The study by Van Ommeren et al. (2000) extends Gronberg and Reed (1994) by allowing job attributes to vary during job spells.

⁴ The comparative statistics of this model have been derived elsewhere (Mortenson,

The hazard rate of quitting h , the instantaneous rate of leaving the job per unit time, is defined as the job arrival rate times the probability of accepting a job offer. Suppose that one observes the hazard rate h , the worker's job attribute x and the worker's wage w . Gronberg and Reed (1994) have shown that the worker's MWP for job attribute x (MWP_x) can be written as:

$$MWP_x = \frac{\partial h}{\partial x} / \frac{\partial h}{\partial w}. \quad (1)$$

So, the marginal willingness to pay for attribute x equals the ratio of the marginal hazard rate of job attribute x to the marginal hazard rate of the wage. This is a useful result. It enables one to estimate workers' willingness to pay for job attributes given duration data on job quitting behavior. It is also intuitively appealing, since the effects of wage and non-wage characteristics on the job hazard rates are informative on the relative importance of the job characteristics for the worker, and therefore on the willingness to pay for a non-wage characteristic.

Job hazard rate models are based on job hazard rates, whereas discrete choice models of job moving behavior are based on the probability of moving job during a fixed interval. Since the probability of moving job can be written as a function of the job hazard rate, it may be expected that MWP estimates can be obtained using discrete choice models. Indeed, it can be shown that equation (1) implies (see Appendix 1 for a formal demonstration):

$$MWP_x = \frac{\partial P}{\partial x} / \frac{\partial P}{\partial w}, \quad (2)$$

1986).

where P is the probability of a worker quitting a job in a certain small interval. Hence, the marginal willingness to pay for a job attribute equals the ratio of the marginal probability of quitting a job of the job attribute to the marginal probability of quitting a job of the wage. Equation (2) enables one to estimate the marginal willingness to pay for nonwage job attributes given an analysis of job mobility using discrete choice models.

2.2. Estimation method

In empirical job mobility studies, it is usually assumed that P , the probability of moving job during a small interval, has one of the following specifications:

$$(i) P = f(\beta_x \cdot x + \beta_w w); \quad (ii) P = f(\beta_x \cdot x + \beta_w \cdot \ln(w)).$$

where β_x and β_w are parameters and $\ln(w)$ is the logarithm of the wage. In both cases, f is a function determined by the choice of the type of discrete choice model (e.g., logit, probit). Using equation (2), specifications (i) and (ii) imply the following formulae for the marginal willingness to pay for attribute x :

$$(i) \quad MWP_x = \frac{\beta_x}{\beta_w}; \quad (ii) \quad MWP_x = \frac{\beta_x}{\beta_w} w.$$

Hence, given consistent estimates of β_x and β_w , the marginal willingness to pay for a nonwage job attribute can be calculated.

2.2.1 Estimation of the standard error of the MWP estimates

The standard error of the ratio of two estimated coefficients can be calculated using the

delta method (see, for example, Goldberger, 1991). The standard error of the estimated MWP is then calculated as $[\text{Var}(\beta_x) + (\beta_x/\beta_w)^2 \cdot \text{Var}(\beta_w) - 2(\beta_x/\beta_w)\text{Cov}(\beta_x, \beta_w)]^{0.5}/\beta_w$, where Var denotes the variance and Cov denotes the covariance. The studies of Viscusi (1980), Robinson (1990) and Gruber and Madrian (1990) do not report the value of $\text{Cov}(\beta_x, \beta_w)$, so we report standard errors fixing the correlation between β_x and β_w to zero. Presuming other values for the correlation, we find that the standard errors hardly change (we will also report standard errors presuming that the correlation equals 0.2).⁵ In appendix 2, we calculate the relative bias in the standard error, which turns out to be small even for relatively high values of the correlation between β_x and β_w . For example, when the correlation is 0.2, then the relative bias is maximally 10%. Finally, for the application in section 3.4, we are able to calculate the difference between estimates based on estimates of $\text{Cov}(\beta_x, \beta_w)$ and estimates based on the assumption that $\text{Cov}(\beta_x, \beta_w)$ equals zero. The difference turns out to be negligible. Hence, the overall conclusion is that, in general, the value of $\text{Cov}(\beta_x, \beta_w)$ hardly influences estimates of the standard error of the estimated MWP, and the assumption that $\text{Cov}(\beta_x, \beta_w)$ is equal to zero is acceptable when calculating the standard error of the estimated MWP using the delta method.

3. Empirical applications

In this section, we derive MWP estimates based on three studies that have examined

⁵ Fixing the correlations to zero is not unreasonable, since we know from the hedonic wage literature that the correlation between job attributes and wages is low (see Brown,

workers’ job quitting behavior in the US: Viscusi (1980), Robinson (1990) and Gruber and Madrian (1994). See Table 1. In these studies, job quitting behavior is analyzed by means of discrete choice models. Moreover, in these studies, the wage and at least one nonwage job attribute are included as explanatory variables of quitting behavior. These studies report that the wage has a negative effect on job quitting behavior. Further, we estimate the employees’ marginal willingness to pay for commuting using a probit model. These results are compared to the same estimate based on a hazard model (Van Ommeren et al. 2000).

Insert table 1

To facilitate comparison of the results, the estimated MWP relative to the mean wage (multiplied by 100) is also reported, denoted as’ MWP(R). One advantage of this measure is that it is the closest to the empirical specifications employed by Robinson (1990) and Gruber and Madrian (1994) in which the wage is logarithmically expressed (see 2.2). In the current study, when interpreting estimates, we use a 5% significance level, except when otherwise stated.

3.1. Viscusi (1980)

Viscusi (1980) focuses on quit behavior of individuals in the 1976 University of Michigan Panel Study of Income Dynamics (PSID). A logit model is employed for males and females separately. A job nonwage attribute variable that measures the industry’s average

1980).

injury and illness rate (INJRATE) is included. This variable is obtained from an external data source. The annual mean of INJRATE is 10.46 (males) and 7.04 (females), respectively. The mean wage rate is 4.40 (males) and 2.85 (females), respectively. The results can be found in Table 2.

Insert table 2

The results show that the males' MWP for INJRATE is -0.106. This is equivalent to 2.37% of the males' mean wage level. This result implies that industries where the risk was 10 percent above average must raise earnings by 2.48%⁶ to prevent male workers from quitting to industries where the risk was average. It must be noted that the estimate is statistically insignificant.

The females' MWP for INJRATE is statistically significant and equal to -0.129. The standard error is 0.047.⁷ This is equivalent to 4.53% of the females' mean wage level. This result implies that industries where the risk was 10 percent above average must raise earnings by 3.18% to prevent female workers from quitting. This estimate is somewhat higher than those derived from conventional hedonic wage estimates, which are usually

⁶ This percentage is obtained by multiplying the males' average INJRATE (10.46) and the estimated MWP relative to the mean wage (2.37), and taking 10% of the attained value.

⁷ The standard error is calculated assuming that the correlation between the estimated coefficients of the wage and INJRATE is zero. When the correlation is assumed to be 0.2, the standard error is equal to 0.049, thus only marginally higher than the value of 0.047 used in Table 2.

less than 2% (see Sider, 1985; Biddle and Zarkin, 1992; Viscusi 1993). Based on the estimates for INJRATE, the implied value of life estimates is in the range of 4.5 • 7 million dollars, which is also higher than most conventional estimates which fall in a range of 2 • 5 million dollars (Viscusi, 1993).⁸

In conclusion, in line with the theoretical and empirical studies cited in the introduction, we find that the MWP estimates for the absence of risk using the empirical results of Viscusi (1980) point to higher estimates than conventional estimates.

3.2. Robinson (1990)

Robinson (1990) analyses the 1978-1980 National Longitudinal Surveys (NLS) of young men and women to study the effect of cancer risks on job quitting behavior. Robinson (1990) uses external data sources to create an index of cancer risk for 231 occupations. Each respondent to the NLS is ascribed his or her occupation's risk of cancer. Different logit models are employed using this index of cancer risk and a self-reported measure for the cancer risk. The earnings are specified in logarithms.⁹ Furthermore, three other types of job attributes are included in the analysis: the absence of job training, of promotion prospects and of job security.

⁸ To calculate the implied value of life estimates, we have used additional information on the relationship between INJRATE and the likelihood of death as provided by Viscusi (1979).

⁹ To facilitate interpretation, we have standardized the original cancer index (by dividing the original cancer index by its maximum value) so that the index has values ranging from zero to one inclusive.

Robinson (1990) shows that the constructed cancer index, a self-reported hazard exposure and the absence of promotion opportunities are positively associated with quitting. Higher wage rates are less likely to quit. Job training and job security do not appear to affect quitting behavior (see Table 3).

Although the effect of the constructed cancer index and the self-reported hazard exposure on job mobility are both significant (at a 10% significance level), only the estimate for the MWP for the constructed cancer index of risk is significant (at a 10% significance level) and equal to 78.2%.

It appears from Table 3 that the MWP estimates for job training and job security are small, but the MWP for promotion prospects is about 65% to 70% of the mean wage (see columns 3 and 5).¹⁰ We will explain by application of search theory that such an estimate for the MWP estimates for promotion prospects is plausible. To simplify matters, we consider a simplified search model that allows us to obtain an explicit solution for the MWP for promotion prospects.

here table 3

Suppose a worker earns wage w and expects to be promoted at rate λ . Promoted workers receive a wage w^p forever ($w^p > w$). The worker discounts the future at rate ρ .

¹⁰ The standard errors are 27.5 and 30.1 respectively, presuming that the correlation between 'log wage' and 'no promotion prospects' is zero. Presuming that the correlation is 0.2, the standard errors are 28.8 and 31.3 respectively, and therefore only marginally higher than those used in Table 3.

Job-to-job mobility is ignored. Lifetime utility V can then be written as $(\rho w + \lambda w^p)/\rho(\rho + \lambda)$. The MWP for λ equals $(w^p - w)/(\rho + \lambda)$ (since $\partial V/\partial w = 1/(\rho + \lambda)$ and $\partial V/\partial \lambda = (w^p - w)/(\rho + \lambda)^2$). So, the MWP for λ is positive, and decreasing and concave in λ .

In Robinson's (1990) empirical specification of job search behavior, a dummy for 'no promotion prospects' is included in the model. The *marginal* willingness to pay for this dummy can be interpreted as the willingness to pay for a certain *level* of λ . The willingness to pay for λ is defined as $[V(\lambda) - V(0)]/\partial V/\partial w$ and can be written as $\lambda/\rho \cdot (w^p - w)$.¹¹

Suppose now that the presence of promotion prospects implies that λ/ρ is 2.00 and promoted workers receive a wage rise of 35%. The latter estimate is based on a study of internal job mobility which shows that the median percentage difference in mean pay across levels is about 35% (van Gasteren, 1999). The promotion rate λ might be 0.10 (McCue, 1996) and the yearly discount rate ρ is 0.05. These assumptions imply that the MWP for promotion prospects is 70% of the mean wage, which is close to the percentage implied by Robinson's empirical study. Of course, other numerical assumptions would have led to different outcomes, but this example indicates that the value of the empirical estimate is not implausible and in line with theoretical predictions.

¹¹ In this example, job-to-job mobility is neglected. Thus the willingness to pay for λ is overestimated. In the case that jobs are offered at rate p and wage offers equal w^p , the willingness to pay for λ can be written as $\lambda \cdot (w^p - w)/(\rho + p)$, which is smaller than $\lambda \cdot (w^p - w)/\rho$.

3.3 Gruber and Madrian (1994)

Gruber and Madrian (1994) study a policy of limited insurance portability that has been adopted by a number of states and the federal government in the US over the past 20 years. Insurance portability has received considerable attention in the health insurance literature in the US (Monheit and Cooper, 1994). The data they use come from the panels of the Survey of Income and Program Participation. A probit model of job turnover is employed. Wages are logarithmically specified. One job attribute is the length of continuation health insurance coverage. Continuation coverage grants individuals the right to continue purchasing health insurance through their *former* employers for a specified period *after* leaving their jobs. The option to purchase health insurance through the former employers has a number of advantages to workers; it is cheaper, not medically underwritten and is more generous than individual commercial coverage. The right to continue purchasing health coverage is, therefore, invaluable when the employee is laid off or moves to another employer that does not offer health insurance.

Job leavers gain more from continuation coverage than those who stay. Let us focus on the extremes. For job leavers who are unlikely to leave the job, the value of the option of one additional month of continuation coverage must be close to zero. On the other hand, for job leavers who anticipate being laid off within a month, the value of this option may be worth hundreds of dollars. For the average worker, the value of this option must be substantially less.

Gruber and Madrian (1994) estimate the effect of the length of continuation

coverage on job mobility of men. The estimated coefficients of the logarithm of the hourly wage and the length of continuation coverage (measured in months) can be found in Table 4. The MWP for a month of continuation coverage is 3.44% of the mean wage. The standard error of this estimate is 0.82%.¹² To calculate the MWP we will focus on a worker who receives the mean hourly wage of 9 dollars, works 180 hours a month and, therefore, earns 1620 dollars per month. The absolute MWP for a month of continuation coverage for this worker is 55 dollars. The MWP for a month of continuation coverage is between 29 and 81 dollars with 95 percent certainty.¹³ So, the overall evidence is that the economic value of the option to continue purchasing health insurance through the former employers is substantial.

Insert table 4

3.4 Marginal willingness to pay for commuting

Recently, Van Ommeren et al. (2000) estimated the employees' willingness to pay for commuting distance based on an analysis of job durations using a job hazard model.¹⁴

¹² This estimate is based on a zero correlation between the log hourly wage and month of continuation coverage. Presuming that the correlation is 0.2, the standard error is 0.85, only marginally higher than the one used in Table 4.

¹³ Gruber and Madrian (1994) provide some specification checks by estimating four other specifications. Based on the estimates of these specifications, the MWP for a month of continuation coverage is estimated between 2.44 and 4.70% of the mean wage.

¹⁴ Van den Berg and Gorter (1997) estimated the marginal willingness to pay for commuting time by an analysis of subjective responses concerning the minimum

Here, we will estimate the employees' MWP for commuting distance based on a probit analysis of voluntary job moves using the same data set. This analysis aims to show that MWP estimates based on discrete choice models and duration models are similar.

The data set used here (called Telepanel), collected in 1992-1993, includes the complete life cycle pattern of Dutch respondents. This data has been used numerous times, so our description here will be brief (for a more extensive description see Van Ommeren, 2000). The data were collected retrospectively. We use a sample of 372 males who were observed during a period of seven years. Here, the main objective is to estimate the MWP for commuting, a topic which has attracted much attention in the field of transport economics (see Van Ommeren et al. 2000; Small, 1992).

The data set allows us to distinguish between voluntary and involuntary job moves due to firing. The data include information on the workers' municipalities of residence and workplace, which allows us to approximate the commuting distance by the distance between the centres of the municipalities. The mean commuting distance is 20 kilometres. About 40% of the persons work in the same municipality as they live. For these persons, we have fixed the commuting distance at one kilometre.¹⁵ We include further a large number of explanatory variables as suggested by economic theory (for a

acceptable wage. This analysis was restricted to unemployed individuals.

¹⁵ To test whether this way of measuring commuting distance affects our results, we have re-estimated the probit model, including a dummy for those who work in the same municipality as they live. This dummy appears to be statistically insignificant. However, as anticipated, the coefficients of commuting distance indicate then a stronger relationship between commuting distance and mobility than reported in the current paper.

full discussion and descriptive data, see Van Ommeren et al. 1999).¹⁶ We investigated the functional forms of the wage and commuting distance by estimating separate probit models with linear and log-linear transformations of the wage and distance. With respect to commuting distance, there is no difference between the results of the two specifications. The results below are based on a log-linear specification of commuting distance. With respect to the wage, a log-linear transformation fits the data better than a linear one (for a similar conclusion, see Van Ommeren et al., 1999). Estimation of a standard probit model is straightforward. However, in the case of panel data, one should preferably allow for individual specific heterogeneity. In the empirical application, we use both approaches. The results can be found in Table 5 (the full results are found in Van Ommeren et al. 1995). It appears that both approaches render almost identical results.

Insert table 5

The MWP for commuting distance are evaluated at an average distance of 20 kilometres and a wage of 20 Guilders per hour, -0.058 and -0.060 Guilders per kilometre.

¹⁷ The MWP estimates of Van Ommeren et al. (2000), based on a range of job hazard

¹⁶ We include six levels of education, number of subordinates, age, calendar year, size of workplace, industrial sector, job tenure.

¹⁷ The reported standard error of the MWP for commuting distance is 0.0279. When calculating the standard error of the MWP, presuming that the correlation of the estimated coefficients is zero, the standard error is 0.0276. This clearly shows that the standard error

models, are -0.051, -0.058 and -0.068, respectively. Thus, the results are essentially the same.

Presuming a working day of 8 hours, the average willingness to pay for one additional kilometre of commuting is -0.46, so employees are willing to pay 2.3% of their hourly wage to avoid one additional kilometre of commuting. The MWP for commuting distance implies a MWP for commuting time, which depends on average commuting speed. Figures on average commuting speed are absent for the Netherlands, but a reasonable estimate is 20-30 kilometres per hour.¹⁸ Presuming an average commuting speed of 25 kilometres per hour, the estimated marginal value of commuting time is 57 percent of the hourly wage rate. This estimate is in line with Small (1992) who concluded from a large number of studies that the average value of time for the journey-to-work trip is estimated to be 50 percent of the hourly wage rate.

4. Conclusion

The current study adapts Gronberg and Reed's (1994) dynamic approach to valuation of job attributes, by analysis of job mobility using discrete choice models. In line with previous studies that relied on dynamic approaches, we find that MWP estimates are higher than conventional static estimates. For example, based on the study of Viscusi (1980) conducted in the United States, we find that the implied value of life estimates are about 4.5-7 million dollars, which is higher than most conventional estimates which fall in

of the MWP is insensitive to the value of the correlation of the estimated coefficients.

¹⁸ This estimate is based on external sources that reveal that one third of the population commutes by bicycle or by foot; car drivers that commute for less than 16 kilometres have an average speed of 32 kilometres per hour (Van Ommeren, 2000).

a range of 2-5 million dollars.

Gronberg and Reed's (1994) valuation approach and the approach used here both assume that workers have identical utility functions, while assuming that jobs are search goods in a dynamic environment.¹⁹ In contrast, the more conventional hedonic wage approach does not assume identical utility functions, but assumes a static environment. A priori, it is impossible to claim that one of the approaches is in general superior (Herzog and Schlottmann 1990; Gronberg and Reed 1994; Van Ommeren et al. 2000). As emphasized by Gronberg and Reed (1994), more research is needed to understand the differences in static and dynamic approaches to valuation of job attributes (for comparison of these approaches, see Gronberg and Reed, 1994; Van Ommeren et al., 2000). We hope that much-needed research on the relationship between static and dynamic approaches to the valuation of job attributes will benefit from the present paper.

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¹⁹ The assumption of identical utility functions may be inappropriate, potentially invalidating MWP estimates. For example, if preferences for job attributes are heterogeneous, then employees are not randomly drawn from the population as a whole with respect to these job attributes (Kiefer, 1988). As a result, job searchers who have a stronger distaste for certain job attributes will be less likely to accept jobs that offer more of these job attributes, and estimates of the value of job attributes will be biased downwards.

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Appendix 1: Estimating the marginal willingness to pay for job a attribute using discrete choice models

The close relationship between discrete choice models and continuous time hazard models is well known (see Meyer (1990)). Denote T as the employment spell. Then the hazard rate $h(t)$ is defined by the equation:

$$\lim_{\varepsilon \rightarrow 0^+} \frac{\Pr(t + \varepsilon > T \geq t)}{\varepsilon} = h(t) \quad (3)$$

The relationship between P , the probability that an employment spell ends before time $t+1$ given that it has lasted until time t , and the hazard rate function $h(t)$, is then the following:

$$P = 1 - \Pr(T \geq t+1 | T \geq t) = 1 - \exp\left[-\int_t^{t+1} h(s)ds\right] = 1 - \exp[-h(t)] \quad (4)$$

given that $h(t)$ is constant between t and $t+1$. Hence, (4) implies

$$\frac{\partial P}{\partial x} / \frac{\partial P}{\partial w} = \frac{\partial h}{\partial x} / \frac{\partial h}{\partial w} \quad (5)$$

So, given (1) and (5), equation (2) follows.

Appendix 2: The relative bias in the standard error of ratio (β_x / β_w)

We demonstrate here that the relative bias, **defined** as the ratio of the difference between the reported standard error (which is based on the assumption that the correlation between the estimated coefficients β_x and β_w equals zero) and the correctly calculated standard error to the reported standard error standard is less than $\frac{1}{2} |cor|$, where cor denotes the correlation between the estimated coefficients β_x and β_w .

Proof

The ratio of the correctly calculated s.e. to the incorrectly calculated s.e. (presuming cor = 0) can be written using the delta method as (where ratio denotes β_x / β_w):

$$\begin{aligned} & \sqrt{\frac{\text{var}(\beta_x) + \text{var}(\beta_w)ratio^2 - 2.ratio.cor.se(\beta_x).se(\beta_w)}{\text{var}(\beta_x) + \text{var}(\beta_w).ratio^2}} \\ & \leq \sqrt{\frac{\text{var}(\beta_x) + \text{var}(\beta_w)ratio^2 + 2.ratio.|cor|.se(\beta_x).se(\beta_w)}{\text{var}(\beta_x) + \text{var}(\beta_w).ratio^2}} \\ & \leq \sqrt{1 + \frac{2.ratio.|cor|.se(\beta_x).se(\beta_w)}{\text{var}(\beta_x) + \text{var}(\beta_w)(ratio)^2}} \end{aligned}$$

$$\leq \sqrt{1+|cor|},$$

since the maximum of $\frac{2 \cdot ratio \cdot |cor| \cdot se(\beta_x) \cdot se(\beta_w)}{\text{var}(\beta_x) + \text{var}(\beta_w)(ratio)^2}$ is obtained when

$$se(\beta_x) = se(\beta_w) \cdot ratio.$$

Finally note that $\sqrt{1+|cor|} < 1 + \frac{1}{2}|cor|$.

Table 1. Studies of job quitting behavior.

Study	Sample	Nonwage job attributes
Viscusi (1980)	1975-1976, men and women	injury rate
Robinson (1990)	1978-1980,14-24 years	hazard exposure, job training, promotion prospects, job security
Gruber and Madrian (1994)	1984-1987, men, 20-54 years	continuation of health benefit coverage when leaving job
Current study	1985-1991, men	Commuting distance

Table 2. Coefficients of quit behavior with respect to attributes, males and females, 1975 (based on Viscusi, 1980) and the marginal willingness to pay for job attributes.

variables	Males			Females		
	coefficient	MWP	MWP(R)	coefficient	MWP	MWP(R)
wage rate	-0.226			-0.389		
		(0.028)			(0.037)	
INJRATE	0.024	-0.106	-2.37	0.050	-0.129	4.53
	(0.020)	(0.121)	(2.71)	(0.018)	(0.047)	(1.65)

Note: Standard **errors** in parentheses. INJRATE is measured as the incidence **rate** per million hours worked. The annual **mean** of INJRATE is 10.46 (**males**) and 7.04 (females) respectively. **MWP(R)** is calculated at the **mean** wage **rate**, which are 4.48 (**males**) and good **strategic** 2.85 (females), respectively.

Table 3: Coefficients of quit decision with respect to job attributes, young men, 1978-1980, and young women, 1980-1982, US, (based on Robinson, 1990) and the marginal willingness to pay for job attributes.

variables	coefficient	MWP(R)	coefficient	MWP(R)
Log wage	-0.715 (0.193)		-0.663 (0.189)	
Cancer index	0.559 (0.304)	-78.2 (43.1)		
Self-reported hazard exposure			0.292 (0.173)	-44.0 (30.8)
No job training	-0.230 (0.177)	32.2 (27.5)	-0.239 (0.177)	36.0 (30.1)
No promotion prospects	0.465 (0.152)	-65.0 (30.2)	0.463 (0.152)	-69.8 (32.8)
No job security	-0.055 (0.171)	7.7 (27.0)	-0.081 (0.171)	12.2 (28.7)

Note: standard errors in parentheses.

Table 4. Coefficients of job turnover with respect to job attributes, men, 1983-1989, US, (based on Gruber and Madrian, 1994) and the marginal willingness to pay for job attributes.

variables	coefficient	MWP	MWP(R)
log hourly wage	-0.134 (0.008)		
month of continuation	0.005	55	3.44
coverage	(0.001)	(13)	(0.82)

Note: standard errors in parentheses.

Table 5. Empirical estimates of discrete choice coefficients for voluntary job-to-job mobility

variables	<u>Cross Section</u>			<u>Panel Data</u>		
	coefficient	MWP	MWP(R)	coefficient	MWP	MWP(R)
wage rate	-0.82 (0.12)			-0.78 (0.12)		
commuting distance	0.048 (0.022)	-0.058 (0.028)	-2.31 (1.14)	0.048 (0.022)	• 0.060 (0.029)	-2.43 (1.20)

Note: staudard errors in parentheses. MWP(R) is based on a working day of eight hours.